# EMPOWERMENT ZONES AND THE HOUSING MARKET IN PARIS INNER CITY

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***Abstract*** - *We test for the impact of local enterprise zones on house prices in inner Paris. In 1997, a new urban policy program was launched in France which created three kinds of Enterprise Zones (ZUS, ZRU and ZFU). Notably, 9 first tier EZs (ZUS) were created in Paris. We use this natural experiment and compare the evolution of flat prices in these areas to that in a control group of neighboring areas (that could also have been converted into ZUS). We develop a new semi-parametric spatial matching methodology. Each sale observed before the reform is matched with a similar post-reform counterfactual. This matching procedure is done for housing units within the EZ, as well as for units in the vicinity of the EZ. In contrast to what have been evidenced in other administrative districts of the Paris Region and for other kinds of EZs (i.e., ZFUs), we detect no significant effect of this expensive EZ program implementation on housing values.*

***Key-words*** - ENTERPRISE ZONES ; HOUSE PRICES ; SEMI-PARAMETRIC SPATIAL MATCHING METHODOLOGY

***JEL Classification*** - C21, R31, R38

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## 1. INTRODUCTION

Empowerment Zone programs consist of tax incentives in areas with a high unemployment level and a low density of locally-based firms. Tax deductions are granted to firms setting up in these areas and creating jobs locally. Following the first experiments in the US (in the 1980s) or in the UK and France (in the 1990s), an important body of empirical literature has tested whether these programs have reached their objectives in terms of jobs or businesses creations and provided low evidence of significant direct effects stemming from locally-targeted EZ programs. However, the purpose of these EZ programs was often not limited to the local labor market. In France, the policies were extended to subsidies for local home ownership and social housing landlords, as well as to the development of public amenities. In the US, the federal EZ program launched in 1994 also involved subsidies for social services and local amenities. These measures may thus have impacted the local housing market.

Our goal is to extend the literature on the magnitude of EZ effect on local home prices. In the US, the number of studies focussing on the real estate consequences of these locally-targeted programs is limited (see e.g., Engberg and Greenbaum, 1999, Busso and Kline, 2007, or Krupka and Noonan, 2009). Most of them provided evidence of a significantly positive, but moderate, effect on property values. To our knowledge, the only study in France is a companion paper (Gregoir and Maury, 2014) putting the emphasis on a very specific French administrative unit (the Seine-Saint-Denis *département*). We extend this analysis and focus on the inner Paris housing market. We use an integrated econometric setup, simultaneously assessing the possible effects of the local program within the boundaries of the EZ and beyond the limits of the EZ. We rely on a a street-by-street mapping of EZs. We estimate the spatial spillover effect with a semi-parametric distance-to-the-EZ-boundary gradient function. Consequently, we can test for the significance of the EZ impact on house price levels within targeted zones, as well as on house price gaps between units within EZs and units in the vicinity of EZs, controlling for the distance between these two units.

In France, the EZ programs were initiated at the end of 1996 (the Round I program hereafter). Three types of targeted zones were defined with a growing amount of tax exemptions: 750 ZUS (*Zones Urbaines Sensibles*), 416 ZRU (*Zones de Redynamisation Urbaines*), selected among the 750 ZUS) and 44 ZFU (*Zones Franches Urbaines*) were selected and geographically-delimited among the most distressed ZUS. The concentration of targeted zones is relatively high in inner Paris, with 9 newly created ZUS. We collect housing sales values, precise location and attributes in 1996 and 1997 (i.e., the year before and the year after the program implementation). We select all pre-reform (1996) sales observations within or in the vicinity of a ZUS. For each observation (the focal point), we construct a post-reform counterfactual (after the Round I implementation), by selecting a fixed number of sales at the same location or an immediate proximity. We then construct the pre- and post- reform differences of the logarithms of house prices for each focal point within and beyond ZUS borders. We proceed in the same manner for a selected control group of zones that did not become ZUSs in 1996, but could have become so, according to the public criteria based on some descriptive statistics. The log of price gaps (for treated ZUS or controls, within or beyond the considered zone) are pooled and a partially linear single-index model is estimated. Among the linear set of covariates, we include the difference of the main house attributes of the reference and counterfactual observations. The non-parametric component (i.e., the single index) includes all other factors that may potentially interact in a non-linear manner: in particular, the way the distance to the closest EZ (or control) boundary affects the log of house price differences is assessed with a non-linear link function $\eta(.)$ modeled with a penalized spline (P-spline) (Yu and Ruppert (2002)). This type of spline offers an interesting trade-off between the flexibility of the functional form and the computational costs.

We detect no significant effect of the EZ program on housing values in Paris, at least in the short run. This result is at odds with the companion contribution, Gregoir and Maury (2014), which detected a significantly negative effect of the EZ program in the Seine-Saint-Denis district. However, this study was based on a different period (Round II implementation in 2004) and another kind on EZ (i.e., second tier EZs converted into third tier EZs).

The rest of the paper is organized as follows. We provide a survey of the literature on empowerment zones in France in the second section. In Section 3, we present the French EZ program and detail its implementation in Paris. We describe our dataset in the fourth section. Section 5 provides our detailed econometric methodology: the spatial matching process, control group selection, the penalized spline estimation of the partially linear single index model. Section 6 presents the results. The last section concludes.

## 2. DESCRIPTION OF THE FRENCH PROGRAM

In the beginning of 1997, the French authorities initiated the first Empowerment Zone program following experiments in the UK and the US. As said above, three categories of EZ have been created with a growing amount of public aid: ZUS (first tier), ZRU (second tier) selected among the ZUS, and ZFU (third tier) selected among the most distressed ZUS. First tier EZs (i.e., ZUSs that are not ZRUs or ZFUs) may theoretically be exempt from local business and property taxes. It should be noted that such public grants are at the discretion of local authorities and are not compensated by the central government. Moreover, local authorities (municipalities) with a first tier EZ must establish a local housing program as well as promote urban regeneration projects. Households wishing to become homeowners in ZUSs benefit under certain conditions from substantial increases in the maximum amount of publicly funded zero-interest rate loans (*Prêt à Taux Zéro*, PTZ) they are entitled to. Hence, even though aid for ZUSs is mainly focused on business creation and local employment (which can indirectly affect the housing market), aid may also impact on local urban projects (housing programs, local amenities) through enlarged incentives to conduct such projects for local authorities which benefit from this support. Real estate values appear to be an ideal measure for assessing the effect of local ZUSs, since the wealth created by these programs should be capitalized in the local housing stock.

Many descriptive studies have been published on the evaluation of different rounds of the French EZ program with contrasting conclusions (see Thélot, 2004, Gilli, 2006, or Ernst, 2008). The two most significant statistical contributions are Gobillon, Magnac and Selod (2010) and Givord, Rathelot and Sillard (2011). Gobillon et al. (2010) propose an assessment of the effects of the first round of the French EZ program. They use a two-stage procedure to estimate the impact of EZs on individual unemployment durations. First, they estimate a stratified partial likelihood estimator with a panel dataset at the municipality level, from 1993 to 2003. They obtain average municipal effects and illustrate the changes in these effects over time. Comparing these results for treated municipalities (those including an EZ) and counterfactual municipalities with similar characteristics, they provide evidence of moderate but significant effects on the French program in the short run (i.e., for the semesters immediately following the policy implementation). No significant effect is detected over larger horizons. Givord et al. (2011) evaluate the consequences of the Round II French EZ program on economic activity: business creations and employment. They use establishment level data from 2002 to 2007 with a geographic mapping methodology, in order to establish the exact street address. They investigate the impact of the program on treated areas (firms located with EZs boundaries) with a two stage method (propensity score to control the probability for an area becoming a ZFU, plus a kernel matching method). They find significant effects of the program on both business creation and employment. It should be noted that they also find evidence of negative spatial spillover effects on areas in the vicinity of EZs.

## 3. DATA

Data on sales prices come from the Paris Chamber of Notaries. We restrict our sample to transactions in existing apartments. New properties and house sales both represent only a small share of the total sales for the Paris Region and their price and physical attributes strongly differ from those of the existing apartments. We collect individual information on the sale price, the floor area of the apartment, the floor level, the construction period of the building, the presence of an elevator in the building, the number of rooms, of bathrooms, of garages, etc. We also have the precise location (street address) of each housing unit. This is of crucial importance for the geographic mapping procedure of the data (i.e., the distinction of dwellings inside or beyond the EZ's boundaries and distance to the closest EZ boundary). Data are available for the 1996-2007 time period, on a monthly basis.

We focus on sales occurring in the inner Paris *département*, one of the eight administrative units of the Paris Region. Nine ZUS have been created in inner Paris in January 1997 (see Figure 1). Table 1 summarizes some descriptive statistics of the main characteristics of people living in each of these nine zones. We compare these results to those obtained with a set of 8 control zones located in the same area of inner Paris than the 9 treated ZUS (i.e., the North East of Paris). These control zones have been spatially delimited by another urban program (CUCS, *Contrat Urbains de Cohésion Sociale*) that has been launched much later (in 2007) to implement joint actions (contract between the state and local governments) to improve the daily lives of residents in neighborhoods experiencing difficulties (unemployment, violence, etc.). Since they were created according to a different program, these control zones did not receive the same public grants as ZUS in 1997. However, Table 1 shows that treated and control zones are not so different when considering the share of low income households, the share of households getting unemployment benefits or the activity rate. Nevertheless, it should be noted that the share of tenants in the social housing sector is much higher in some ZUS than in control zones (98.1% in the *HBM Aubervilliers* EZ and 99.1% in the *Porte de Montmartre* EZ). In theory, the modelling should be developed to take into account for the possible bias implied by the endogeneity of the Round I selection process. In a similar framework, such a modelling did not allow the econometricians to exhibit the impact of a particular variable on the dependent one in the selection process (see Gregoir and Maury, 2014).

**Figure 1. Map of Inner Paris**


*Grey areas are Round I ZUSs (treated zones).*
*Black areas are zones of the control group (CUCS).*

**Table 1. Descriptive statistics on treated and control zones in Paris**

| Name | low income | social housing | unemp. benefits | activity rate | employees/ workers |
|---|---|---|---|---|---|
| HBM Aubervilliers (T) | 25.0 | 98.1 | 19.7 | 70.0 | 78.2 |
| HBM Ménilmontant (T) | 14.5 | 46.1 | 17.6 | 75.4 | 46.1 |
| La Goutte d'Or (T) | 25.5 | 18.6 | 22.1 | 74.8 | 61.5 |
| Porte Saint-Denis (T) | 11.2 | 6.4 | 18.2 | 82.1 | 38.2 |
| Fontaine au Roi (T) | 17.5 | 14.9 | 19.7 | 79.8 | 47.5 |
| Porte de Saint-Ouen (T) | 15.6 | 61.5 | 17.5 | 77.7 | 64.0 |
| Porte de Montmartre (T) | 23.0 | 99.1 | 19.4 | 69.2 | 78.8 |
| Curial (T) | 16.8 | 59.1 | 20.4 | 77.0 | 63.1 |
| Belleville (T) | 22.6 | 34.4 | 22.0 | 77.6 | 54.7 |
| Buisson Saint-Louis (C) | 15.4 | 22.5 | 20.1 | 80.0 | 50.8 |
| La Chapelle hors ZUS (C) | 19.3 | 20.9 | 19.3 | 75.7 | 61.5 |
| Amiraux Simplon (C) | 15.1 | 14.1 | 18.7 | 78.0 | 55.6 |
| Curial hors ZUS (C) | 21.2 | 47.3 | 19.5 | 75.6 | 60.9 |
| Amandiers hors ZUS (C) | 15.5 | 37.9 | 21.0 | 78.8 | 48.8 |
| Porte de Saint-Denis hors ZUS (C) | 11.7 | 7.1 | 18.5 | 82.5 | 40.5 |
| Porte de Montmartre hors ZUS (C) | 20.0 | 80.4 | 19.7 | 71.5 | 69.9 |
| Porte de Saint-Ouen hors ZUS (C) | 15.6 | 63.0 | 17.9 | 77.6 | 64.5 |
| Paris Metropolitan Area | 8.8 | 21.6 | 15.2 | 81.1 | 49.3 |

*9 treated (T) and 8 control (C) are considered. 'Low Income' is the share of low income households in 2009. 'Social Housing' is the share of tenants in a social dwelling in 2009. 'Unemp. benefits' is the share of households where at least one member get unemployment benefits. 'Activity rate' is the activity rate among 25-64y in 2009. 'Emp/workers' is the share of employees or workers in 2007. Source: INSEE.*

## 4. METHODOLOGY

Let $p_{i,c,t}$ be the log price of individual housing unit *i*, within (or in the vicinity of) EZ *c*, at time *t*. We suppose that this outcome is generated by a semi-parametric model of the form:

$$p_{i,c,t} = X'_{i,c,t}\beta + \mu\left(T_c, I_{i,c}, D_{i,c}, Z_{i,c,t}\right) + \theta_{c,i} + \varepsilon_{i,c,t} \qquad \text{(D.1)}$$

where $T_c$ is a treatment dummy variable (indicating whether or not the EZ *c* was treated in the Round I program in 1997). Hence, $T_c = 1$ is for the observations in the treated group and $T_c = 0$ in the control group. $I_{i,c}$ is a dummy variable indicating if housing unit *i* is located within (treated or the control) EZ *c*, and $D_{i,c}$ is the distance (as the crow flies) between housing unit *i* and the closest EZ border ( $D_{i,c} = 0$ if $I_{i,c} = 1$ ). $X_{i,c,t}$ and $Z_{i,c,t}$ both are vectors containing information on the housing unit's characteristics, the EZ's characteristics and house market trends. We further discuss which variables should be included in the linear part of the model (i.e., included in vector $X$ as done in traditional house price hedonic models) and which ones should be included as arguments of the unspecified function $\mu(.)$. $\beta$ is the hedonic parameter vector. $\theta_{c,i}$ is a spatial location fixed effect. It captures unobserved local neighborhood characteristics that may potentially be correlated with the treatment variable $T_c$ or the variables summarizing the distance to the nearest EZ ($I_{i,c}$ and $D_{i,c}$).

The model is a partially linear semi-parametric model that encompasses the standard linear hedonic setup $X'_{i,c,t}\beta$. The link function $\mu(.)$ is not specified *ex ante*. This semi-parametric specification allows for a non-linear impact on house prices of some house attributes $Z_{i,c,t}$, or the distance to the nearest EZ $\lfloor I_{i,c}, D_{i,c} \rfloor$, as well as their interaction with the treatment variable $T_c$. For instance, we expect home prices to be positively related to $D_{i,c}$, but dependent on some of the apartment's attributes, its size for example: the immediate proximity of an EZ may influence the price differently for large family dwellings (households with children) than for small apartments. We then include the dwelling's floor area variable among $Z_{i,c,t}$ instead of $X_{i,c,t}$. Furthermore, we do not know how the distance gradient may evolve after the new policy implementation. Consequently, the interaction between $T_c$ and $D_{i,c}$ should not be specified ex ante (both are arguments of $\mu(.)$).

The estimation procedure is as follows. We first collect all pre-reform housing sales $p_{i,c,t}$ ($t < 1997$). Let $n$ denote the total number of such observations. To control for individual and local unobserved heterogeneity, each reference sale *i* is spatially matched with a fixed number ($\bar{n}$) of post-reform housing sales $\bar{p}_{j,i,c,t'}$ ($t' \geq 1997$, $j = 1,...,\bar{n}$). Different values for the number of nearest neighbors are tested. The matching procedure is done according to the sole spatial criterion (i.e. we gather the $\bar{n}$ post-reform sales nearest to the pre-reform reference point $i$). We then compute the housing price differential between transaction $i$ and the counterfactual transactions *j*, i.e. $\Delta p_{i,j,c,t,t'} = \bar{p}_{j,c,t'} - p_{i,c,t}$. All these $n \times \bar{n}$ observations are pooled and weighted according to the distance between the focal and counterfactual sales. Let $w_{i,j}$ denote the weight between sales *i* and *j*. $\bar{n}^{-1} \sum_{j=1,...,\bar{n}} w_{i,j} = 1$ for each *i*. In our framework, each pair $(i, j)$ shares the same fixed unobserved effect $\theta_{c,i}$. This is reasonable if the distance between *i* and *j* remains low. Accordingly, the housing price differential becomes

$$\begin{aligned}\Delta p_{i,j,c,t,t'} = & \left(X'_{j,c,t'} - X'_{i,c,t}\right)\beta \\ & + \left[\mu\left(T_c, I_{j,c}, D_{j,c}, Z_{j,c,t'}\right) - \mu\left(T_c, I_{i,c}, D_{i,c}, Z_{i,c,t}\right)\right] \\ & + \left(\varepsilon_{j,c,t'} - \varepsilon_{i,c,t}\right)\end{aligned} \tag{D.2}$$

where we take advantage of the fact that $I_{j,c} = I_{i,c}$ (if the focal point is within an EZ's boundaries, the spatial matching procedure is done inside the EZ, no counterfactual sale is selected beyond the EZ boundaries and similarly if the focal point is outside the EZ) and $D_{j,c} - D_{i,c} \approx 0$ to reformulate the nonparametric part of the model (second term in bracket in the RHS of preceding equation)

$$\begin{aligned}\bar{\mu}\left(T_c, I_{i,c}, D_{i,c}, Z_{i,c,t}, Z_{j,c,t'}\right) \equiv & \ \mu\left(T_c, I_{j,c}, D_{j,c}, Z_{j,c,t'}\right) \\ & - \mu\left(T_c, I_{i,c}, D_{i,c}, Z_{i,c,t}\right)\end{aligned} \tag{D.3}$$

Such a general setup is subject to the curse of dimensionality due to the large number of covariates (arguments of function $\bar{\mu}$). We limit our approach to a partially linear single-index. The non-linear link function is then defined as $\eta\left(\gamma Y_{i,j,c,t,t'}\right) \equiv \bar{\mu}\left(T_c, I_{i,c}, D_{i,c}, Z_{i,c,t}, Z_{j,c,t'}\right)$ where $Y_{i,j,c,t,t'} = \left(T_c, I_{i,c}, D_{i,c}, Z_{i,c,t}, Z_{j,c,t'}\right)$ and $\gamma$ is a parameter vector. This single-index assumption implies a lesser degree of generality, since all interactions of elements of $Y$ are captured by the eventual non-linearity in the univariate function $\eta(.)$, but it may still capture important features of our high-dimensional problem and will greatly facilitate the numerical estimation procedure. Applications of the partially linear single-index model may be found in various fields, including housing economics (see for example, Wang et al. (2007) who propose an application of this model to the case of Boston House prices, using a different setup). We implicitly assume that conditionally on the covariates, the growth in housing prices is not related to the treatment variable, $\Delta p_{i,j,c,t,t'}$ ? $T_c$ | $(X, Z, \eta, I, D)$.

Following Yu and Ruppert (2002), we model $\eta(.)$ with a penalized spline (P-spline hereafter). As detailed by Ruppert and Carroll (1997) or Ruppert (2002), P-splines are a generalization of smoothing splines allowing more flexibility in the choice of knots and penalty parameters. Moreover, according to other methods with different link function specifications, P-splines allow for easier and faster regression procedures. For example, we have also estimated our setup with local linear approximations and the minimum average variance estimation (MAVE, see Xia and Härdle, 2006), but convergence was much less rapid. The P-spline function for $\eta$ is:

$$\eta(a) = \delta_0 + \delta_1 a + \ldots + \delta_p a^p + \sum_{k=1}^{K} \delta_{p+k}\left(a - \kappa_k\right)_+^p \tag{D.4}$$

where $\{\kappa_k\}_{k=1}^{K}$ are the spline knots. They capture the jumps in the $p^{th}$ degree of the polynomial. $\delta = \{\delta_0, \delta_1, ..., \delta_{p+K}\}$. Following Yu and Ruppert (2002), the estimation is carried out with several numbers of knots $K$, ($K = 3,\ 5, 10$, more than $10$ knots is above the usual practice in applications). The knots are equally spaced along the quantiles of the index, $\gamma Y_{i,j,c,t,t'}$. We adopt the standard penalized and weighted least squares estimators of $\beta$, $\gamma$ and $\delta$ which minimize:

$$P_{n,\bar{n},w,\lambda}(\beta,\gamma,\zeta,\delta) = \left(n\bar{n}\right)^{-1} \times \sum_{i=1}^{n}\sum_{j=1}^{\bar{n}} \left\{ w_{i,j}\left[\left(\bar{p}_{j,c,t'} - p_{i,c,t}\right) - \left(X'_{j,c,t'} - X'_{i,c,t}\right)\beta - \eta\left(\gamma Y_{i,j,c,t,t'}\right)\right]\right\}^2 + \lambda\delta' D\delta \tag{D.5}$$

where $D$ is a positive semidefinite symmetric matrix. We select $D$ equal to a diagonal matrix with its last $K$ diagonal elements equal to $1$ and the rest is zero (Ruppert and Carroll (1997)). We penalize the jumps in the P-spline. $\lambda \geq 0$ is the penalty parameter. Its value is chosen by a standard generalized cross-validation (GCV) selection process which is simultaneously run with the non-linear least squares estimation algorithm. The complete procedure (estimation + GCV) follows closely Yu and Ruppert (2002) approach. After selecting the size of the nearest-neighbors kernel ($\bar{n}$), the polynomial degree ($p$), the number of knots ($K$), we obtain initial estimates for $\beta, \zeta$ and $\gamma$ with OLS on a linear (fully parametric) version of the model. A grid of values for the penalty parameter ($\lambda$) is created and $P_{n,\bar{n},w,\lambda}(.)$ is first minimized over $\delta$ (with OLS estimates of $\beta, \zeta$ and $\gamma$) and then over $(\beta, \zeta, \delta, \gamma)$ for each value of $\lambda$. A GCV score is computed for the selection of $\lambda$ (see equation 21 in Yu and Ruppert, 2002) and the estimation outcome with this value is finally chosen.

## 5. RESULTS

We present here the results of our P-spline partially linear single-index model. The chosen sample period is [1996; 1997]: i.e., we consider apartment sales one year before the EZ implementation (1996), and one year after (1997). We select the control group as explained in the data presentation section: 8 zones (CUCS) than have not been converted into ZUSs in 1997, but could have been according to their characteristics. We pool all pre-reform sales observations within or close to treated areas (i.e., ZUSs) and within or close to control group zones: close to is defined here as being within a 300 meter radius of the nearest EZ.

**Table 2. Estimation. Inner Paris for the 1996 to 1997 period**

| *Hedonic Terms:* $X\beta+\eta$ | Estimates | s.e. |
|---|---|---|
| Level 1 (ref = ground level) | $0.0919^{**}$ | 0.0113 |
| Level 2 | $0.1308^{**}$ | 0.0112 |
| Level 3 | $0.1306^{**}$ | 0.0111 |
| Level 4 with elev. | $0.2006^{**}$ | 0.0108 |
| Level 4 no elev | $0.1403^{**}$ | 0.0113 |
| 2 rooms (ref = studio) | $0.0239^{**}$ | 0.0082 |
| 3 rooms | $0.0524^{**}$ | 0.0119 |
| 4 rooms | $0.0564^{**}$ | 0.0160 |
| 5+ rooms | $0.0636^{**}$ | 0.0209 |
| Parking space (ref = no) | $0.0779^{**}$ | 0.0112 |
| Bathroom (ref = no) | $0.1486^{**}$ | 0.0057 |
| constr. $\in [1948,1969]$ (ref<1948) | –0.0255 | 0.0171 |
| constr. $\in [1970,1980]$ | $0.0622^{**}$ | 0.0178 |
| constr. $>$ 1981 | $0.1720^{**}$ | 0.0253 |
| $2^{nd}$ quarter (ref = $1^{st}$ ) | –0.0030 | 0.0071 |
| $3^{rd}$ quarter | –0.0029 | 0.0072 |
| $4^{th}$ quarter | 0.0037 | 0.0067 |
| *Index parameters:* $\gamma$ | Estimates | s.e. |
| $T_c$ | 0.0081 | 0.0357 |
| $I_{i,c}$ | $-0.7201^{**}$ | 0.3139 |
| $D_{i,c}$ | 0.0015 | 0.0023 |
| $T_c \times I_{i,c}$ | 0.6404 | 0.4070 |
| $T_c \times D_{i,c}$ | –0.0030 | 0.0024 |
| $Z_{i,c,t}$ (log floor area reference) | $0.2541^{**}$ | 0.0107 |
| $Z_{j,c,t'}$ (log floor area counterfactual) | –0.0822 | 0.00901 |
| *P-spline parameters:* $\delta$ | Estimates | s.e. |
| $\delta_0$ | $-0.1025^{**}$ | 0.0123 |
| $\delta_1$ | $0.0564^{**}$ | 0.0087 |
| $\delta_2$ (jump) | $0.0890^{**}$ | 0.0085 |
| $\delta_3$ (jump) | 0.0123 | 0.0095 |
| $\delta_4$ (jump) | $0.0233^{**}$ | 0.0092 |
| $\delta_5$ (jump) | $-0.0382^{**}$ | 0.0096 |
| $\delta_6$ (jump) | $-0.0885^{**}$ | 0.0095 |

$\lambda$=0.0025, N=7,659. ** = 5% signif, * = 10% signif.
Standard errors are calculated with block-bootstrap procedures

We set $p = 1$ (first degree P-spline polynomial part), as usually done in the literature. We set $n = 3$ (results are qualitatively similar for other values of $n$). Each pre-reform sale is then spatially matched with 3 post-reform counterfactual sales. The whole sample size is 30,141, which includes all differences in the log of house prices, $\Delta p$, between the pre-and post-reform periods. The benchmark control group chosen is also included.

Vector $X_{i,c,t}$ includes a large set of structural attributes for each dwelling: the floor level and presence of an elevator, the number of rooms, the presence of a parking space, a bathroom, the period of construction of the building, the year and quarter of transaction. The only covariate included in $Z_{i,c,t}$ is the log of floor area (in square meters). This continuous variable allows us to capture interactions between the size of dwellings and the treatment effect. Finally, the list of variables used to constitute the index are: $T_c$, the treatment dummy; $I_{i,c}$ the within-EZ dummy; $D_{i,c}$ the distance (in hundreds of meters) to the nearest EZ boundary; $Z_{i,c,t}$ the log of floor area of the pre-reform reference housing sale $i$; and $Z_{j,c,t'}$ the log of floor area of post-reform counterfactual housing sale $j$.

We first conduct some preliminary estimates (obtained with standard parametric methods), in order to assess the relationship between the endogenous variable – the difference in the log of house prices $\Delta p_{i,c,t,t'}$ – and the univariate index $\delta Y_{i,j,c,t,t'}$. We then select the number of knots $K = 5$. As suggested by Ruppert and Carroll (1997), the total number of knots should not exceed 10. In our case, additional knots do not appear to capture significant new locally-linear trends, in the relationship between the endogenous variable (controlling for the linear hedonic term) and the index.

Table 2 gives estimates of model (differential). Results are reported for the three different sets of parameters: $\beta$, $\delta$, $\zeta$ and $\gamma$. We emphasize that the estimates of the index parameter vector $\gamma$ should not be directly quantitatively interpreted, since they depend on the link function $\eta(.)$ and its parameter estimates. The parameter estimates for the standard hedonic part of the model: i.e., $X\beta$, have the expected sign and magnitude. We put the emphasis on the results for the index $\mu(T_c, I_{i,c}, D_{i,c}, Z_{i,c,t})$. We detect no significant impact of the treatment dummy variable $T_c$ on the dynamics of housing prices between 1996 and 1997. Second, the impact of covariate $(T_c \times I_{i,c})$ on the price growth rate between the pre- and post-reform years is positive, though not significant. The magnitude of the impact of this covariate on home values cannot be directly assessed from the parameters estimates (since it depends on the parameters of the p-spline) and we obtain the ATT (Average Treatment Effect for the Treated) with simulation methods (see Gregoir and Maury, 2014, for a detailed explanation): we find that growth in home prices between 1996 and 1997 has been slightly more than 3 percentage points higher within treated areas, though once again this effect is non significant. Hence, we do not provide evidence of a positive incidence of the French Round I program on targeted EZ in inner Paris. Looking at the data, it seems that this outcome stems from the heterogeneity in home price dynamics between ZUS in Paris: some of them experienced price growth rates above the average of control zones between 1996 and 1997, while other treated zones did not. These differences in price evolution may be the consequence of uncontrolled factors in the benchmark model, but could also be linked to the voluntary nature of aids in first tier EZs (the local authorities can decide in each case to grant aid or not, and these aids will not be compensated by the state). Maybe we would have obtained different results with second or third tier zones (notice that there are only first tier zones in Paris), since tax abatements are mandatory for them. In a companion paper, Gregoir and Maury (2014), working with another administrative unit (the Seine-Saint-Denis département), conclude that the implementation of the Round II local development program in 2004 seems to have contributed negatively to the local value of the housing market, suggesting there has been a stigmatizing effect on areas that became ZFUs (i.e., third tier EZs).

Finally, we detect no significant gap in home prices dynamics between housing units located inside a treated EZ ( $I_{i,c} = 1$ , $T_c = 1$) and housing units located in their immediate vicinity ( $I_{i,c} = 0$ , $T_c = 1$, $D_{i,c} \approx 0$ ). The positive impact of the $(T_c \times I_{i,c})$ covariate is dampened by the negative impact of the $I_{i,c}$ covariate. On the contrary, price have grown less inside a non-treated (control) zone ( $I_{i,c} = 1$, $T_c = 0$ ) than in their immediate vicinity ( $I_{i,c} = 0$ , $T_c = 0$, $D_{i,c} \approx 0$ ). We find no significant relationship between the distance to the closest EZ border $(D_{i,c})$ and the growth rate of property values.

## CONCLUSION

In a flexible semi-parametric framework, we analyze the impact of a particular EZ policy implemented in 1997 on real estate values within and in the vicinity of nine areas in Paris inner city. Whereas such policies are designed to regenerate urban environment through incentives to conduct local amenities projects and financial aid to new homeowners, we do not detect any significant impact of this program on housing prices at any level in comparison with what has been observed in areas, similar from a social and economic descriptive point of view but not included in the program. This is at odd with the objectives of such policies and with the results obtained in a companion study on their impacts on suburb areas in which a significant negative impact was detected. This may be due to the heterogeneity of the selected areas, some may have been positively impacted by the policy and some negatively, or to the fact the policy was undersized.

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## LE MARCHÉ DU LOGEMENT PARISIEN ET LES ZONES URBAINES SENSIBLES

***Résumé*** - *Nous estimons l'impact des zones franches sur le marché du logement parisien. En 1997, trois types de zones franches ont été mis en place (ZUS, ZRU and ZFU). En particulier, 9 zones urbaines sensibles (ZUS) ont été créées dans Paris intra-muros. Nous comparons l'évolution des prix des appartements dans ces zones à celle observée dans des zones voisines comparables. Nous proposons une nouvelle méthode semi-paramétrique avec appariement géographique. Chaque vente observée avant la réforme est couplée à un ensemble de ventes observées après la mise en place de la réforme. Cette procédure d'appariement est effectuée pour des logements situés dans ou à proximité d'une ZUS. Nous ne détectons pas d'effet significatif de ce programme de zones franches sur les valeurs immobilières.*

***Mots-clés*** - ZONES URBAINES SENSIBLES, PRIX DE L'IMMOBILIER, APPARIEMENT GÉOGRAPHIQUE